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ABSTRACT

"Workfare" schemes that offer poor participants unskilled jobs at low wages have become a popular alternative to cash or in-kind handouts. Yet little is known about a key determinant of the cost effectiveness of such schemes in reducing poverty: the behavioral responses through time allocation of participants and their families. These responses affect the foregone incomes of participants and, hence, the net transfer benefits. This paper estimates how time allocation within sampled households responded to new rural employment opportunities under the "Employment Guarantee Scheme" (EGS) of the State of Maharashtra, India. Data were obtained from households surveyed over 6 years in two villages, Shirapur and Kanzara, that differed in wealth, land ownership patterns, agricultural factors, and occupational structure. Participation in the projects was affected by gender, village, and household wealth and caste. Literacy and education had little effect on EGS employment except in Kanzara, where head of household's education was significant and positive. In both villages, EGS employment was generally exogenous to time allocation, suggesting that the ideal of providing work on demand was not being met. Significant gender cross-effects in time allocation were observed, as well as displacement of different activities for different genders: unemployment for men, leisure and domestic activities for women. Overall, the projects generated sizable net income gains to participants, certainly far greater than implied by using market wages to value foregone income. Transfer benefits alone led to a reduction in poverty, of almost the same magnitude as a uniform and undistorting allocation of the same gross budget. Contains 34 references, 13 data tables, and statistical equations. (SV)

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LSMS
Living Standards
Measurement Study
Working Paper No. 100

Income Gains for the Poor from Public Works Employment

Evidence from Two Indian Villages

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Income Gains for the Poor from Public Works Employment

Evidence from Two Indian Villages

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Number 100

Income Gains for the Poor from Public Works Employment

Evidence from Two Indian Villages

Gaurav Datt
Martin Ravallion

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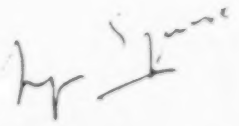
Foreword

In efforts to better target public outlays for poverty reduction, "workfare" schemes have been a popular alternative to cash or kind handouts. Workfare participants are offered unskilled jobs at relatively low wages. Yet we know very little about one of the key determinants of the cost-effectiveness of such schemes in reducing poverty: the behavioral responses through time allocation of the participants and their families. Those responses will determine, in part, the foregone incomes of the participants, and (hence) the net transfer benefits. This paper provides estimates of how time allocation within sampled households responded to new rural employment opportunities provided under the "Employment Guarantee Scheme" of the State of Maharashtra in India.

The paper is one of a series documenting results of a research program in the Poverty and Human Resources Division of the Policy Research Department. The program has aimed to improve our knowledge about the impacts on poverty of some of the targeted poverty alleviation schemes found in developing countries. It is hoped that intensive analysis of a few selected schemes will also yield lessons for the design and assessment of anti-poverty schemes in a wider range of settings.



Lyn Squire
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Abstract

Current knowledge provides little guidance on one of the key issues in evaluating workfare schemes: What are the net income gains to participants? This paper offers an answer for rural public employment in the state of Maharashtra in India. An econometric model of intra-household time allocation is proposed, and the paper offers a consistent estimator, recognizing that the model entails that both regressands and endogenous regressors will be censored. The empirical implementation indicates that workfare projects induce significant behavioral responses, though the predominant time displacement is such that the net income gains remain large. Employment on the projects led to a reduction in poverty, of almost the same magnitude as a uniform and undistorting allocation of the same gross budget.

1 Introduction

Public works projects have been a popular policy instrument for poverty alleviation in developing countries[1]. The net income gains to participating workers will depend in part on how time allocation across activities and persons responds to the new employment opportunities. Impacts on intra-household activity patterns are of interest for other reasons; for example, it has often been asked: when a woman takes up new employment on a development project does this augment, or displace, her domestic labor? Quite generally, the opportunity cost of the labor employed in a new project will depend on behavioral responses through time allocation; that will depend in turn on preferences, and on the way markets and institutions work in determining the constraints facing individuals. Thus the social evaluation of a new project will require evidence on, or assumptions about, behavioral responses through time re-allocation.

One can distinguish two quite different ways of addressing these issues in past analyses:[2]

i) It is often argued that there is substantial "disguised unemployment" in underdeveloped rural economies, such that the opportunity cost of the labor employed in a new project will be low. This approach typically emphasizes quantity constraints on behavioral responses, such as due to the existence of wage inflexibility, or social customs in the gender division of labor.

ii) Alternatively, one can appeal to the neoclassical model of time allocation, in which the assumption of wage and price flexibility guarantees that any allocation consistent with time and other resource endowments is feasible. This model gives less straightforward behavioral predictions, though the opportunity cost of labor will almost certainly not be zero, and may well be quite high. For example, some past estimates of the net income gains from direct poverty reduction schemes have used prevailing market wage rates for similar work.[3]

In each case, the outcome will also depend on whether one is concerned with the opportunity cost in terms of forgone utility (the "welfarist" approach), or that in terms of forgone incomes (a common

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[1] On the arguments for and against this type of anti-poverty policy in developing countries see Drèze and Sen (1989), Drèze (1990), World Bank (1990, Chapter 6), Ravallion (1991a), Besley and Coate (1992), Dev (1992) and Lipton and Ravallion (1993). For a more general discussion of targeted anti-poverty schemes see Besley and Kanbur (1993).

[2] For surveys of the arguments and approaches to measurement see McDiarmid (1977) and Rosenzweig (1988).

[3] See, for example, Hossain's (1988) evaluation of the famous Grameen Bank in Bangladesh, suggesting low net income gains.

"non-welfarist" approach).[4] In principle, this distinction is independent of that between the "disguised unemployment" approach and the neoclassical approach. In practice, however, it is more common to find that those who follow the former approach tend to emphasize income gains and losses, while those following the latter adopt the welfarist approach by which the choice maximand - utility - is the sole metric of well-being.

This paper takes a fresh look at behavioral responses to the new employment provided by public-works projects, and their implications for assessing transfer benefits. The essential idea is to model the intra-household allocation of time *conditional* on existing public-works employment; we call this the conditional time allocation model (CTAM). The theoretical formulation of CTAM is general enough to encompass both approaches described above. The quantification and valuation of displaced time relies largely on the empirical evidence provided by the estimated CTAM. Though our approach could be used to help inform either "welfarist" or "non-welfarist" evaluations, we will apply it here solely to the issue of assessing the net income gains from public employment projects. This will throw empirical light on the arguments that have been made about the efficiency of rural development projects as instruments for the alleviation of income poverty. In the process, we shall also examine the question of whether households appear to be rationed in their access to public employment; the ability of public-works schemes to provide work to whoever wants it at the going wage has long been recognized as a key determinant of their success in alleviating poverty.[5]

We estimate CTAM on household data for two villages in the state of Maharashtra in India, where rural public-works employment is available under that state's "Employment Guarantee Scheme" (EGS).[6] An explicit aim of this scheme is to alleviate rural poverty through the income gains to the participating workers. Unskilled manual labor is provided at low wages (on a par with the agricultural wage). The work is mainly small scale rural public works projects, such as roads, irrigation facilities, and re-forestation. EGS was introduced as a statutory program in the mid 1970s and expanded rapidly to reach average annual attendances of about 100 million person days in recent years. It is financed almost entirely out of taxes on the urban sector of the State of Maharashtra, including an income tax levy. Our data are from a longitudinal household level survey done over many years by well-trained resident

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[4] On the distinction between "welfarist" and "non-welfarist" approaches, see Sen (1979).

[5] For further discussion on this point see Ravallion (1991 a,b).

[6] There is a large literature on this particular scheme, including Dandekar (1983), Acharya (1990), Dev (1992), Bhende et al. (1992), and Ravallion et al. (1993). For other references see Ravallion (1991a).

investigators for the International Crops Research Institute of the Semi-Arid Tropics (ICRISAT), India. The data are widely considered to be of high quality, and they appear to provide a unique opportunity for evaluating household behavioral responses to the EGS.

The following section outlines our model of time allocation in the abstract, and the properties of our estimator. Section 3 introduces the setting and the data. Section 4 presents the non-parametric estimates of foregone incomes, and notes some of their limitations. In section 5, we discuss the parametric CTAM and its estimation. We then go on to present and discuss the results in sections 6 and 7, while section 8 and 9 examine their implications for evaluating the net transfer benefits from the projects, and the impacts on poverty respectively. Our conclusions are summarized in section 10.

2 The Time Allocation Model

In this setting an econometric model of time allocation should be at least consistent with the existence of involuntary unemployment of labor. This rules out formulations which assume that *any* time allocation consistent with the overall time endowment is feasible. We believe that one cannot dismiss either the casual evidence, or that from the survey data we are using, suggesting that quantity constraints on time allocation are common in this setting. For example, Walker and Ryan (1990) quote estimates of the rate of involuntary unemployment in the ICRISAT villages in 1975-76 of 19 percent for men and 23 percent for women.[7] This aggregate figure hides a good deal of variation; the rate of unemployment in slack periods is far higher, being 39 percent for men and 50 percent for women. Insufficient wage employment is clearly the main factor here. But a number of constraints on time allocation (both internal and external to the household) appear to underlie such figures. For example, there appear to be binding constraints on what sort of farm work different genders can do, as a result of which men will often be idle in one season, while women are clearly over-worked. Our model should be consistent with a potentially complex web of constraints on time allocation.

The model predicts time devoted to an activity (wage labor, own farm work and other self-employed work, leisure and domestic work, and unemployment) as a function of exogenous variables *and* the amount of time spent on public-works projects. Consider the household's choice problem of allocating total available time $\bar{L}$ across n activities, where the last activity is public-works. Income is

$$y = F(L_1,..,L_n,x) \tag{1}$$

where x is a vector of relevant variables (prices, fixed factors in production and taste parameters). The vector x can be taken to include $\bar{L}$. The chosen allocation maximizes

$$U(y,L_1,..,L_n,x) \tag{2}$$

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[7] These are on a daily status basis, giving the number of days in which work was desired but not obtained as a proportion of the total number of days of work desired.

In addition to the time constraint, $\Sigma L_i = \bar{L}$, an optimum must satisfy a set of rationing constraints

$$L_i \in [0, L_i^{max}] \tag{3}$$

where it is assumed that the maximum time that can be devoted to activity i is itself a function of the household or individual characteristics, as

$$L_i^{max} = \Phi_i(x) \tag{4}$$

This last assumption can be justified in a number of ways. For example, it may reflect work-related disability. Or it may reflect social determinants of how employment is rationed, possibly to preserve some tacit collusion to keep wages above market clearing levels.[8]

The optimal allocation of time across the n activities then maximizes

$$U[F(L_1,..,L_n,x),L_1,..,L_n,x] + \lambda_0(\bar{L} - \Sigma L_i) + \Sigma\lambda_i(L_i^{max} - L_i) \tag{5}$$

where the λ_i (i=1,..,n) are non-negative Lagrange multipliers. We can write the solution as

$$\begin{aligned} L_i &= L_i(x, L_n) \quad i=1,..n-1 \\ L_n &= L_n(x) \end{aligned} \tag{6}$$

in which we make explicit the conditionality of L_i (*for* $i \neq n$) on L_n. We refer to the functions $L_i(x, L_n)$ as the "*conditional time allocation model* (CTAM)". Note that one can always write solutions in this form, whether or not the allocation to any activity i is rationed. To do so, one solves the first-order and complementary slackness conditions for activities i=1,..,n-1 conditional on the (rationed or un-rationed) solution for L_n. This does not assume that L_n is exogenous.

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[8] On such models of wage determination in this setting see Datt (1989) and Osmani (1991).

3 Setting and Data

Our data are for households surveyed over six years (1979-80 to 1984-85) in two Maharashtra villages, Shirapur and Kanzara.[9] The villages differ in a number of respects. Kanzara is relatively a richer village, has more assured rainfall, is better irrigated and has less participation in the EGS. The villages also differ in terms of the cropping pattern, the occupational structure, and the distribution of land holdings. These differences have implications for household time allocation pattern which we discuss later.

The data allow us to distinguish the time allocation of each person across five activities:

1: wage labor other than on the public works projects (mainly agricultural labor)

2: own farm labor and labor on handicrafts/trade activities

3: unemployment (number of days for which work was sought but could not be found)

4: leisure/domestic work

5: wage labor on public works (mainly EGS)

These categories are self-explanatory, though we would make two observations:

i) Unfortunately, the data do not allow leisure to be differentiated from domestic work. For women, a large proportion is probably domestic labor, which may also include some farm activities.

ii) The fact that positive unemployment is reported (as noted in section 2) suggests that one or more activities must be rationed. It also indicates that the EGS is falling short of its aim of providing work to whoever wants it.[10]

Tables 1 and 2 give cross-tabulations of days of participation against various variables, including time allocation by gender, for Shirapur and Kanzara respectively. Some points worth noting about the results in Tables 1 and 2 include:

i) There is no sign of a displacement of wage labor time for females in either village, though there is such displacement for males, particularly in Kanzara. However, the propensity to do wage work of all forms will tend to be correlated with other variables such as landlessness and caste; later we will see whether any displacement is evident when one controls for these other variables.

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[9] For details on the agro-economic profile of these villages, see Bhende (1983) and Kshirsagar (1983). Also see Walker and Ryan (1990) for further information on the socio-economic conditions in these and other ICRISAT villages.

[10] There is supportive evidence for that conjecture from an independent source; see Ravallion et al. (1993).

ii) For both villages and genders, there is a stronger suggestion of displacement in own-farm labor, though it should also be noted those who are employed most in public works tend to own the least land to farm; this is evident in the (rapid) decline of both LANDI and LANDU as public-works participation increases (bottom panels of Tables 1 and 2). We will see if the effect survives when one controls for land-holding.

iii) Where we do see stronger, and more convincing, signs of displacement is in leisure/domestic work, which declines steadily as public-works employment increases.

4 Estimating Foregone Income: A Non-Parametric Approach

As a first approximation, we used a simple non-parametric approach. The time allocation of non-participants in EGS is used to predict what the time allocation of participants would have been had they not worked on EGS. A simple comparison of the average time allocation of all non-participants with that of all participants does not control for other differences in the characteristics of participating and non-participating households which also influence their time allocations (Tables 1 and 2). A better approach would be to make this comparison conditional on some important determinants of household time allocation. For example, we may compare the time allocation of land-poor participating households with that of non-participating households who are also land-poor. Since such comparisons involve partitioning the sample by the conditioning variables, there are obvious limits to how far this approach can be taken given the overall sample size.

The sample of 33 households in each village is first partitioned into three equal-sized groups on the basis of their average real asset holdings over the 6-year period, which is possibly the single most important determinant of household time allocation.[11] Each asset group is further partitioned into participating and non-participating sub-groups. Let $s_k^{m,g}$ denote the average share of activity k in total available time for adult male members of all *non-participating* households in asset group g.[12] The female share $s_k^{f,g}$ is defined analogously. Then, for a *participating* household i in year t, the time displaced in different activities may be estimated as

$$D_{kit}^{j,g} = s_k^{j,g} T_{it}^{j,g} - L_{kit}^{j,g} \qquad (for\ j = m,f)$$

where $T_{it}^{j,g}$ is the total available time for adult members of gender j in household i in year t.[13] Displaced time is estimated separately for each village. For Kanzara, male and female displaced time were not computed separately owing to the limited public-works participation of women.

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[11] See Tables 2 and 3. Other variables - such as land owned - matter, but are correlated with real assets.

[12] The average share is defined as the mean, over all non-participating household years in any asset group, of the ratio (L_{kit}^{m}/T_{it}^{m}), where the T_{it}^{m} is the total available days for adult male members in household i in year t, of which L_{kit}^{m} are devoted to activity k.

[13] Note that, by construction, displaced time in all activities will add up to the time allocated to public works.

We do not expect this procedure to yield credible estimates of displaced time for *every* participating household-year; the variation in time allocation amongst participants is too great to allow that. However, average displacement may be better tracked. Thus, in Table 3 we report the average displaced time in different activities over all participating household years (across all asset groups). Table 3 also gives our valuations of the average displaced time.

The general principle is to value displaced time at household-year specific prices. With the non-parametric estimates, since we will be concerned only with average displaced time (rather than its distribution), the valuation is done at average "prices", where the average is constructed over all household-years with participation. The displaced wage labor time is valued at household- and gender-specific average daily wage rates for any given year. For the valuation of time displaced in own farm and other self-employed activities, we use the normalized quadratic profit function estimated for these villages in Datt (1989). This profit function is estimated at the village level using data for all the ten villages, including Shirapur and Kanzara, surveyed by ICRISAT. The arguments of the profit function include gross cropped un-irrigated area, gross cropped irrigated area, family labor, owned bullock labor, the daily wage rate for hired labor, rental rate for hired bullock labor, and a composite price for seeds, fertilizers, pesticides, manures and machine hours. The profit function is used to derive a village and year specific estimate of the marginal profit per acre with respect to family labor time, which is then assumed to apply to both own farm and other self-employment activities. The foregone income in these activities is then calculated by multiplying this marginal profit per acre by the household's gross cropped area and the time displaced in these activities. The important point is to use a *marginal* and not *average* valuation. The prevailing agricultural wage rates may grossly overstate the value of marginal time displaced in own-farm and other self-employment activities. Time displaced from unemployment and leisure/domestic work is assumed to entail zero foregone income.

The estimates in Table 3 suggest that foregone income represented 30 percent of earnings from public works in Shirapur (43 percent for males, 10 percent for females) and 4 percent in Kanzara. Some aspects of the estimates in Table 3 are, however, rather implausible. The estimates for males in Shirapur indicate that, as against their mean public-works employment of 58 days, 80 days of leisure/domestic work are displaced, and their unemployment would have been significantly *higher* in the absence of public works. Clearly, what is producing these results is that the households participating in public works are also the ones with high rates of unemployment and labor-force participation; partitioning by real assets is not a adequate control. The problem seems less serious for females in Shirapur and for Kanzara. But, probably in all cases, displaced time in leisure/domestic work is over-stated and that in unemployment is under-stated.

5 The Econometric Model and Estimation

The problem with the above approach is insufficient control for the determinants of time allocation. A parametric model holds greater promise from this point of view. Following the discussion in section 2, the general (gender-differentiated) CTAM can be written as follows. Let L_{kit}^{j} denote time allocation to activity k by persons of gender $j(=m,f)$ in household i in time period t. The equations for time allocation across all other activities are:

$$L_{kit}^{m*} = \beta_k^{m'} x_{it} + \gamma_k^{mm} L_{nit}^{m} + \gamma_k^{mf} L_{nit}^{f} + u_{kit}^{m} \tag{7}$$

$$L_{kit}^{f*} = \beta_k^{f'} x_{it} + \gamma_k^{fm} L_{nit}^{m} + \gamma_k^{ff} L_{nit}^{f} + u_{kit}^{f} \tag{8}$$

(for $k=1,n-1$; $i=1,H$; $t=1,T$) where x_{it} is a vector of explanatory variables, and

$$L_{kit}^{j} = L_{kit}^{j*} \quad \text{if } L_{kit}^{j*} > 0$$

$$L_{kit}^{j} = 0 \quad \textit{otherwise}$$

for j=m,f. The dependent variables are of course censored at the lower bound of zero.

The time allocation equations are specified for three of the four non-EGS activities: wage labor, own farm labor (including other self-employment), and unemployment. Leisure/domestic work is determined residually, though we discuss the implications of relaxing this. However, since the days worked on any of these activities are censored variables, there is no obvious way of imposing additivity in the form of cross-equation parameter restrictions (Pudney, 1989).

Estimation of the CTAM poses a difficult econometric problem, in that the (limited dependent variable) model contains a censored right hand side variable (L_n) which may well be endogenous. In Appendix 1, we derive a consistent estimator for the CTAM, and an exogeneity test for the censored right hand side variable. This estimator is a generalization of that proposed by Smith and Blundell (1986) who allow only continuous endogenous variables in the limited dependent variable model. Following that approach, our estimation strategy is outlined below.

First, the model (7)-(8) is written conditional on error processes v_{nit}^{m} and v_{nit}^{f} as

$$L_{kit}^{m*} = \beta_k^{m'} x_{it} + \gamma_k^{mm} L_{nit}^{m} + \gamma_k^{mf} L_{nit}^{f} + \alpha_k^{mm} v_{nit}^{m} + \alpha_k^{mf} v_{nit}^{f} + \epsilon_{kit}^{m} \tag{9}$$

$$L_{ikt}^{f*} = \beta_k^{f\prime} x_{it} + \gamma_k^{fm} L_{ikt}^{m} + \gamma_k^{ff} L_{ikt}^{f} + \alpha_k^{fm} v_{ikt}^{m} + \alpha_k^{ff} v_{ikt}^{f} + \epsilon_{ikt}^{f} \tag{10}$$

where the equations for EGS employment are

$$\begin{aligned} L_{ikt}^{m*} &= \pi^{m\prime} x_{it} + v_{ikt}^{m} \\ L_{ikt}^{f*} &= \pi^{f\prime} x_{it} + v_{ikt}^{f} \\ \text{where } L_{ikt}^{j} &= L_{ikt}^{j*} \text{ if } L_{ikt}^{j*} > 0, \quad L_{ikt}^{j} = 0 \quad \text{otherwise} \end{aligned} \tag{11}$$

Model (9)-(10) is estimated by the Tobit maximum likelihood (ML) estimator after replacing v_{ikt}^{m} and v_{ikt}^{f} by their consistent estimates; the latter are obtained as residuals from equations (11) for male and female employment on EGS respectively. Exogeneity of male and female public employment to the time allocation for any activity is then tested by the significance of the corresponding α parameter. For any time allocation equation, exogeneity is tested sequentially for male and female public employment, beginning with the α parameter with the lower t-ratio.

If the hypothesis of exogeneity of male or female employment on the public works projects is found to be statistically acceptable (we use a conservative significance level of 10 percent), then the time allocation model is re-estimated assuming exogeneity. If both male and female public employment variables are found to be exogenous, we are left with the standard tobit model. In this case, we further prune down the model to exclude either of the public employment variables if they turn out to be highly insignificant (with t-ratios less then unity).

Next, the time displaced in any activity k due to male and female participation in public works (denoted D_{ikt}^{m} and D_{ikt}^{f}) is estimated as the difference between the expected days of work in activity k with no participation in public works and the expected days of work in that activity conditional on the household's current level of participation. Thus

$$\begin{aligned} D_{ikt}^{j} = {} & E\left[L_{ikt}^{j} \mid x_{it}, L_{ikt}^{m}=0, L_{ikt}^{f}=0, v_{ikt}^{m}=0, v_{ikt}^{f}=0 \right] \\ & - E\left[L_{ikt}^{j} \mid x_{it}, L_{ikt}^{m}, L_{ikt}^{f}, v_{ikt}^{m}, v_{ikt}^{f} \right] \quad (\text{for } j = m, f) \end{aligned} \tag{12}$$

The model is estimated using longitudinal data of six years duration for the two villages with 33 households each.[14] For Shirapur, we shall estimate the CTAM separately for male and female

[14] Despite the panel structure of the data set, we were unable to exploit it for estimating a fixed (or random) effects time allocation model because of the censored nature of the dependent variables. Apart from the inconsistency of a fixed effects Tobit estimator for a short panel, to estimate a fixed effects tobit model we would have had to exclude all households with zero participation in any activity for all six years in the panel (Heckman

household members, allowing for gender differences in time allocation, with cross-effects across genders (so, for example, the female time spent working on the own-farm may increase when the male joins the public works project.)

A gender disaggregation was not feasible for Kanzara owing to the very limited participation of women in the EGS there, resulting in very few non-limit observations (see Table 2). This made both the parameter estimates of the female public employment equation and its residuals (included amongst the regressor variables of the CTAM) sensitive to minor changes in the specification. We therefore opted for aggregating male and female time allocations for Kanzara. For similar reasons, the model is not separately estimated for children. Given their extremely low participation in the labor force and still lower participation in the projects, we assume that their public-works employment comes entirely out of the time they were not in the labor force.

The x_u vector of explanatory variables in the time allocation model includes the variables listed in Table 4. Apart from the year dummies, five sets of variables are included, viz. variables relating to (i) household size and composition, (ii) value and composition of household assets, (iii) caste and educational status, (iv) incidence of some form of work disability, and (v) total available time (days per year) for adult males and females. The last set of variables allow for the overall time constraint (section 2), and are constructed as the total reporting days (the number of days for which the respondent provided time allocation information) *minus* days of sickness or non-residence in the village. The wage rates for both public and private work are deliberately excluded because of their potential endogeneity. EGS employment is remunerated on a piece-rate basis, and thus the time wage rate is not independent of the level of employment. The wage rate for private employment also has an endogeneity problem since the average wage received by a household member is an employment-weighted average of wage rates for different agricultural and non-agricultural operations performed by him or her over the year.

Note that the x_u vector of explanatory variables used in the public-works employment equations is taken to be the same as that in the conditional time allocation model. Thus there is a potential identification problem. The problem is less of a concern for non-linear models, of which the CTAM is an example. In particular, the usual exclusion restrictions are typically not required for identifying non-linear simultaneous models (Amemiya, 1985). For the CTAM, identification is possible by exploiting

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and MaCurdy 1980). For our data set, this would have meant throwing out more than half the observations in many cases. A further potential problem would be that the effective sample would have varied enormously across activities and gender. However, we do attempt to capture household-specific effects on time allocation by including in the set of regressors variables which are invariant over time for a household, for example, the caste ranking of the household, 6-year average of real assets of the household.

the non-linearity of the Tobit predictions from the public employment equations. That is our approach here, although we did test the robustness of our results to this choice by also estimating a model where the x_u vector in the public employment equations had two additional variables, quadratic terms in CASTE and SCHYRH (see Table 4). The results for the latter (not reported, but available from the authors) were very similar to the case with identical set of exogenous variables in the public employment and conditional time allocation equations. For Shirapur, the estimates of time displaced by EGS employment in all activities except unemployment were identical in the two cases; for Kanzara too, the estimates of displaced time were quite similar in the two cases.

6 Discussion of the Results

Table 5 gives the estimated parameters of the models determining employment on the public works projects. The salient features of these results are noted below.

i) Male participation in the projects (in Shirapur, where the gender break-down is feasible) fluctuates far more over time than does that of females, as is evident from the year dummy variables in Table 5. The year 1982-83 was the year with highest village mean income in Shirapur, and this may well explain the drop in male employment in 1982-83, with workers being attracted into other activities. The drop in the following year is not explicable the same way; that year was no better than average. In Kanzara, we also see a sharp fall in employment on EGS in 1983-84, which is also the year in which average income peaked in that village, though the coefficient is barely significant at the 10 percent level. Of course, one must also consider the possibility that we are observing the effects of some form of rationing of EGS employment, such as through the opening and closing of works in progress; this appears to be an important factor in the EGS, at least over recent years (Ravallion, Datt and Chaudhuri, 1993).

ii) The effect of long-term wealth (MRAST being the six-year mean of real wealth) on EGS employment differs between genders and villages. At the mean point, wealthier households tend to work less on these projects (holding the other variables constant) in Shirapur, though the relationship has different curvature for males and females (convex in the former case, concave in the latter). The effect is also concave in Kanzara, though the coefficients are not significantly different from zero.

iii) Ceteris paribus, higher caste status is also associated with reduced public works employment in Shirapur, though not Kanzara. For example, in Shirapur, low wealth but high caste households do less of that work than do otherwise identical low caste households. Thus, some form of caste-related social stigma seems to be in operation here.

iv) There is little to suggest that literacy or education affects employment on the projects, except in Kanzara, where SCHYRH is significant and positive. Again it should be noted that this effect holds constant other variables, including wealth. So the positive sign in Kanzara may be picking up a direct effect of education on the likelihood of participating in public employment, through (for example) knowledge of one's rights under the EGS.

The effect of land owned on EGS employment is also quite different between genders and villages. In Kanzara, households with more un-irrigated land tended to work less on the projects, but there is no significant effect of irrigated land. In Shirapur, household land-holding has no significant effect on EGS employment of females, while (at the mean points) the effect is positive on male

employment. Of course, since wealth is held constant, it is not clear that the latter result implies poor targeting to landless households. It is also worth noting that the turning point in the relationship between male EGS employment in Shirapur and un-irrigated land owned is quite close to the mean, so that the relationship is positively sloped amongst those with relatively low land, and negatively sloped amongst larger holdings.

vi) The only significant effects of livestock ownership are in the female equation for Shirapur, and only for livestock other than bullocks. The effect is negative at the mean.

vii) The demographic variables come out quite strongly in Shirapur, but not Kanzara. As one would expect, larger households in Shirapur tend to have larger employment in the projects, ceteris paribus.

viii) Work disability tends to reduce EGS employment, though the effect is only strongly significant amongst females in Shirapur.

ix) Total available days (excluding days sick, or absent from the village) have the expected positive sign in each gender's equation in Shirapur, though there is no indication of significant gender cross-effects. The effect is not evident in Kanzara, though possibly the need to combine genders is confounding it there.

x) As a general comment though, we do note that the equations for Shirapur appear to be better estimated than that for Kanzara; the possibility that we may not have a particularly good set of instrumental variables for public employment in Kanzara should be kept in mind in interpreting the rest of our results.

Of greater interest are the estimates of the CTAM which are given in Tables 6 to 8. The salient features are as follows:

i) The tests for exogeneity of EGS employment are indicated by the estimated α-parameters and their t-ratios. We find that for both genders and both villages, exogeneity of EGS employment is accepted for both wage labor (other than public works) and own farm labor (including trade and handicraft activities), but not unemployment. Thus only for the unemployment equation do we retain the residuals from the first stage Tobit models for public-works employment in Table 5. The endogeneity of EGS employment in the unemployment model, but not other equations, could reflect the fact that employment on EGS requires a minimum spell of work (typicaliy two weeks). Those experiencing greater unemployment are then more likely to obtain whatever EGS employment is available.

ii) The conclusion that the public works employment is exogenous to other activities suggests that such employment is rationed. This is consistent with the reported unemployment, and also with the

findings (using aggregate time series data) of Ravallion, Datt and Chaudhuri (1992).[15] Failure to obtain this work whenever needed will tend to undermine the social insurance function of public-works schemes, and it may also facilitate the possibilities for corruption.

iii) For males in Shirapur, higher wealth is (ceteris paribus) associated with higher wage labor supply and lower unemployment at the mean point, though there is no significant effect on own farm activities. For females, higher wealth implies lower wage labor supply as well as lower unemployment, while there is a positive effect on own farm labor time, though it could barely be considered significant statistically. In Kanzara, wealth has a negative effect on wage labor supply, but no significant effects on other activities.

iv) Ceteris paribus, high caste households are less likely to do wage labor for both genders and villages, though the effect is only strongly significant for males in Shirapur. High caste is associated with lower male unemployment in Shirapur. There are no significant effects of caste on own farm labor time in Shirapur, but some sign of a positive effect in Kanzara.

v) Literacy tends to have a positive impact on recorded unemployment in Shirapur for both males and (though less significant) females; there is little to suggest effects on other activities. The sign of the effect of literacy on unemployment is reversed in Kanzara, where there are also indications of quite pronounced effects on other activities, with a shift out of wage labor time into own farm and other activities as the number of literate persons in the household increases. However, more years of schooling for the household head is associated with the reverse switch across activities in Kanzara. We have no explanation for this pattern.

vi) The effects of land ownership on time allocation are pretty much as one would expect for males in Shirapur (with a switch out of wage labor time into own-farm time), but the same impacts are not evident on female labor time. Livestock assets follow a similar pattern, though positive effects on female own farm time become evident.

vii) Male disability significantly curtails male wage labor time in Shirapur. Female disability does not do the same to female wage labor time. Furthermore, male disability results in higher female wage labor time, ceteris paribus.

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[15] Bhende, Walker, Lieberman and Venketaram (1992) also report that many of the chronically unemployed in these villages, who were not accommodated on EGS sites because of an excess turnout of laborers, became discouraged and did not return to work sites.

7 Specification Tests

We conducted specification tests for non-normality and heteroskedasticity in the public employment equations, using the test proposed by Chesher and Irish (1987).[16] The tests are based on second and higher moment residuals, where the rth moment residual is defined (suppressing sub-scripts) as

$$\hat{e}^{(r)} = d.\hat{E}(\eta^r | d=1) + (1-d).\hat{E}(\eta^r | d=0) - \mu^{(r)}$$

where d is an indicator function with value 1 for non-censored observations, η is the standardized error (v/σ), the operator $\hat{}$ indicates that the expectations are evaluated at the maximum likelihood estimates, and $\mu^{(r)}$ is the rth moment of the standard normal distribution. The non-normality test is motivated by noting that expected value of the third and fourth moment residuals is zero under the maintained hypothesis of normality; the heteroskedasticity test is motivated by noting that the covariance between second moment residuals and a set of exogenous variables is zero under the maintained hypothesis of homoskedasticity. Since we have a large number of exogenous variables in the public employment equations, we instead base the heteroskedasticity test on the squared predicted values of the regression function.

The test results are given in the top panel of Table 9. For Shirapur, the assumption of normality is strongly rejected although in the case of males, this is entirely on account of the rejection of the fourth moment (kurtosis) condition. For Kanzara, while the normality assumption is acceptable at the 5 percent per cent level of significance, both the third and fourth moment conditions are individually rejected. Except for females in Shirapur, the tests also indicate significant heteroskedasticity.

To examine further the sources of misspecification, the second, third and fourth moment residuals are plotted in Figures 1(a), (b), and (c). Under the maintained hypothesis of spherical errors, the expected value of all moment residuals is zero. Figure 1 suggests that the moment residuals depart significantly from zero only for a limited number of observations[17]: a single household for Shirapur males and for Kanzara, and two households in case of Shirapur females. We thus re-estimated the model deleting these households; three in case of Shirapur and one for Kanzara. The specification tests

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[16] These tests are also discussed by Gourieroux et. al. (1987) and Pagan and Vella (1989).

[17] The observations in Figure 1 are sorted by ascending values of 6-year mean real assets of households, and year.

subsequent to the deletion are reported in the bottom panel of Table 9. The test statistics indicate the assumption of normality is now acceptable in all cases; heteroskedasticity is also attenuated in all cases, though still significant for Shirapur males and for Kanzara.

Given our main interest in deriving estimates of displaced time (and hence foregone incomes) associated with public-works participation, we further look at the effect of deleting the 'problem' households on the estimates of the key parameters determining time displacement. The results are given in Table 10, which shows the estimated parameters of public-works employment in the time allocation equations, both for the full and the pruned sample. Two observations can be made. First, it turns out that the results for the exogeneity of public-works employment are unaffected by pruning; as before (see section 6), exogeneity is rejected only for the time spent in unemployment in either village. Second, the parameter estimates are quite similar in the two cases, with only one notable exception for Kanzara, where the pruned sample indicates a significantly lower displacement of wage labor time. In the light of these results, we chose to use the entire sample, recognizing that this may yield an over-estimation of foregone incomes.

Figure 1: Moment Residual Plots

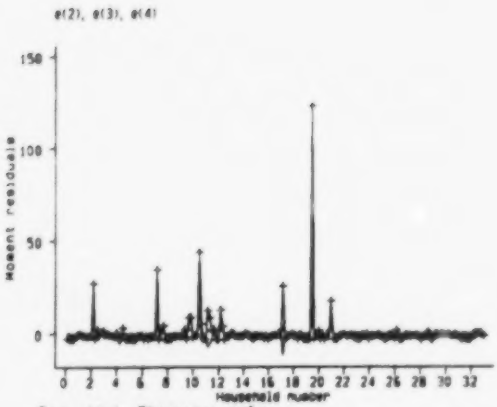

Fig. 1(a): Shirapur male

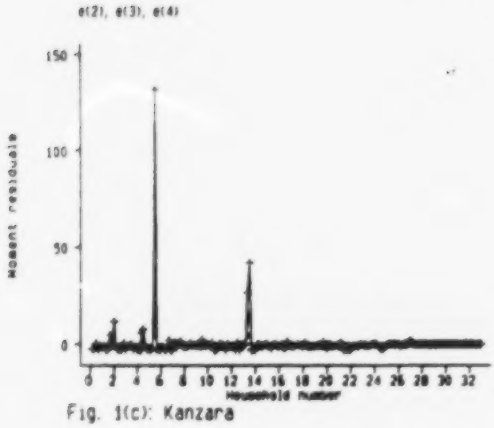

Fig. 1(c): Kanzara

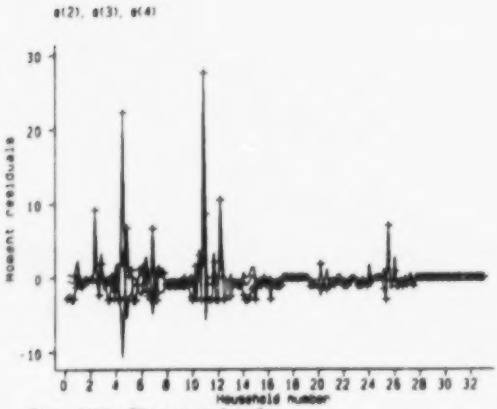

Fig. 1(b): Shirapur female

8 Implications for the Transfer Benefits from Public-Works Projects

We shall only discuss here the implications for the average transfer-benefits; elsewhere we examine the distribution of those benefits, and policy implications (Ravallion and Datt, 1993). Table 11 gives a summary of the implied displacement of labor time in Shirapur for each activity associated with extra time working on EGS sites. The corresponding results for Kanzara are given in Table 12. The following points are of interest:

i) For males in Shirapur, there is a strong negative displacement of time unemployed by extra EGS employment. When averaged over all participants in the EGS, the probability of the non-limit observation of unemployment for males in Shirapur is 0.77 (so that only 23 percent of pooled male household-years are predicted to have zero unemployment). Thus about 60 percent of an extra week of EGS employment came out of unemployment for males in Shirapur. The rest is made up largely of a displacement of own farm labor (Table 11).[18]

ii) For females in Shirapur, there is little sign of significant displacement effects in these three activities, so that extra time in EGS employment comes almost entirely out of leisure or domestic work (Table 11). We also ran the leisure/domestic labor equation separately (rather than retrieving it residually); the estimated marginal displacement was -1.1, close to the figure in Table 11.

iii) There is no displacement of *own* wage labor for both males and females in Shirapur. All of the displacement of wage labor time in this village occurs through the cross-effects.

iv) There is an indication of quite strong cross-effects of female EGS employment on male time allocation, though the reverse effect (of male public employment on female time allocation) is less evident. Male time allocation to own farm labor increases when female employment on EGS increases, while male time allocation to other wage labor falls. These effects are more difficult to interpret. Possibly some of what is being classified as "domestic work" by females in the data set is actually an own farm activity, and this is what men are taking up when the female joins the EGS project. The extra male own farm labor appears to be coming out of other wage labor and leisure.

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[18] We also directly estimated the displacement effects using a separate model for leisure/domestic work (i.e., without imposing additivity). We found that there is a larger displacement of leisure/domestic work for males than implied by the other equations of CTAM. However, the key conclusion on foregone incomes is unaffected, given that marginal foregone income from leisure/domestic work is assumed to be zero.

v) The overall displacement to time allocation from public employment is quite different in Kanzara. The bulk of the time is coming out of wage labor and own farm activities (Table 12). The foregone income is then going to be larger, as we discuss below.

Table 13 gives our estimates of the average number of days in the various activities displaced by public works, and the value of the corresponding income losses. The average is for all participating household-years, defined as those with non-zero adult male or female participation. The basis of valuation is the same as for the non-parametric estimates in section 4, except for one difference, *viz.*, the valuation uses household-year-specific, rather than average, prices.

It turns out that foregone income per displaced day in own farm activities is much lower than that in wage labor. Three factors may be at work here. First, the marginal contribution to farm profits of displaced days in own farm activities is found to be much less than the average profit per day of own farm labor. Second, the marginal contribution of family labor is an increasing function of gross cropped area (Datt, 1989), and thus tends to be low for the bulk of public-works participants who are small operators. Third, the extra time in own farm work may often be spent in tasks, such as soil conservation, whose contribution to farm output and profits, though probably small, is difficult to detect empirically. If true, the last explanation implies some under-estimation of foregone incomes.

For Shirapur, the main activity displaced is unemployment for males and leisure/domestic work (which, unfortunately, cannot be split from the survey data) for females. In the aggregate, slightly over 40 percent of the time spent on the EGS site came out of leisure/domestic work in Shirapur, while one third came out of unemployment. Only one fifth involved a sacrifice of other wage labor time. The pattern is rather different in Kanzara, where nearly a third came out of other wage labor time, and a quarter was from own farm activities. Foregone incomes are estimated at 21 percent of gross wage earnings from public works in Shirapur, and 32 percent in Kanzara. The overall level of EGS employment was lower in Kanzara, though (given the pattern of displacement), its pecuniary opportunity cost was higher.[19] This is intuitive; the same factors which drive up participation would presumably also reduce foregone income. Net transfer benefits from public works generated on average (for participating household-years) a 10 percent increase in pre-transfer earnings in Shirapur, a 7 per cent increase in Kanzara.

Finally, we ask how the assessment of transfer benefits would change if the prevailing market wage is used as the opportunity cost of public-works employment. Using household-year specific (also

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[19] This result supports the similar conjecture in Bhende, Walker, Lieberman and Venketram (1992).

gender-specific in case of Shirapur) average daily agricultural wage rates, it turns out that foregone incomes would be 93 percent of gross earnings from public works in Shirapur (102 percent for males; 86 percent for females), and 77 percent in Kanzara. The differences amongst our various estimators are small compared to this difference; we conclude that the prevailing market wage greatly overestimates the foregone income of participating workers.

9 Impact on Poverty

To assess the distributional impact of net earnings from public employment, Figure 2 gives the actual ("post-intervention") cumulative distribution of six-year mean income and that of the corresponding "pre-intervention" income (actual income *minus* gross workfare earnings *plus* foregone income).[20] To help assess the role played by foregone incomes, we also give the "pre-intervention" distribution that one would have obtained if foregone incomes were assumed to be zero (so this is simply actual income minus gross earnings from EGS).

Two points are notable from Figure 2:

i) Allowing for foregone incomes, earnings from public employment unambiguously reduce poverty; this would hold for any poverty line, and any poverty measure within a broad class.[21] Of course, this does not mean that poverty is lower than it would have been under some alternative method of disbursing the same gross budget; we pursue this question somewhat further below.

ii) The greater effects of allowing for foregone incomes (as measured by the vertical distances between the "pre-intervention" and "zero foregone income" lines in Figure 2) tend to be found at the lower end of the distribution. For example, ignoring foregone incomes one would have concluded that the proportion of the population with an income less than Rs 500 per person per month fell by about seven percentage points (from about 10 percent of the sample population to under 3 percent) as a result of earnings from EGS. However, the impact is a far more modest two percentage points when one allows for foregone incomes.

It is beyond the scope of this paper to properly assess the performance of workfare in reducing poverty relative to alternative policies with the same aim. However, we can preform one simple test, in which the counter-factual allocation is a uniform transfer of the same gross budget across all households in both villages (whether poor or not).[22] We assume that this can be done at negligible administrative cost, and with negligible impact on the pre-intervention income distribution (such as through income effects on time allocation or household dissolution). Neither assumption is realistic, and they will

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[20] The cumulative distribution functions (CDF) in Figure 2 have taken into account the stratified random nature of the sample; the densities at any income level have thus been obtained by using the relevant inverse sampling rates. For clarity in Figure 2 we have truncated the distributions at the top, though this does not alter any of the conclusions that follow.

[21] This follows from the first-order dominance tests for poverty comparisons; see Atkinson (1987).

[22] Note that uniformity across households does not require information on household size.

probably lead us to over-estimate the impact on poverty of uniform transfers.[23] We are probably also under-estimating the impact of the EGS on poverty since we do not consider induced income effects such as through asset creation and the effect on agricultural wages. We also assume that the EGS wage bill accounts for two-thirds of the gross budget; this figure is the estimate obtained by Ravallion, Datt and Chaudhuri (1993) from their analysis of the accounts of the Maharashtra Employment Guarantee Scheme.

Figure 2 also gives the distribution of post-transfer incomes implied by uniform transfers under these assumptions. Though it can be seen that uniform transfers do not first-order dominate the distribution achieved by earnings from public employment, there is only a very narrow range in which the actual distribution function lies below that implied by uniform transfers. And uniform transfers second-order dominate the actual distribution, implying that a wide range of poverty measures indicate a lower poverty under uniform transfers for all poverty lines (Atkinson, 1987).

This last calculation should not be taken to imply that the poor would be better off if one abandoned the rural public works projects in favor of uniform transfers; there are a number of other factors that would need to be considered, as noted above, including the results of specification testing (section 7) which suggest some under-estimation of net income gains from workfare. However, Figure 2 does lead us to question any presumption that the costs of this form of targeting - namely the foregone incomes and the non-wage costs - are of negligible consequence to an assessment of the policy's cost-effectiveness in reducing poverty.

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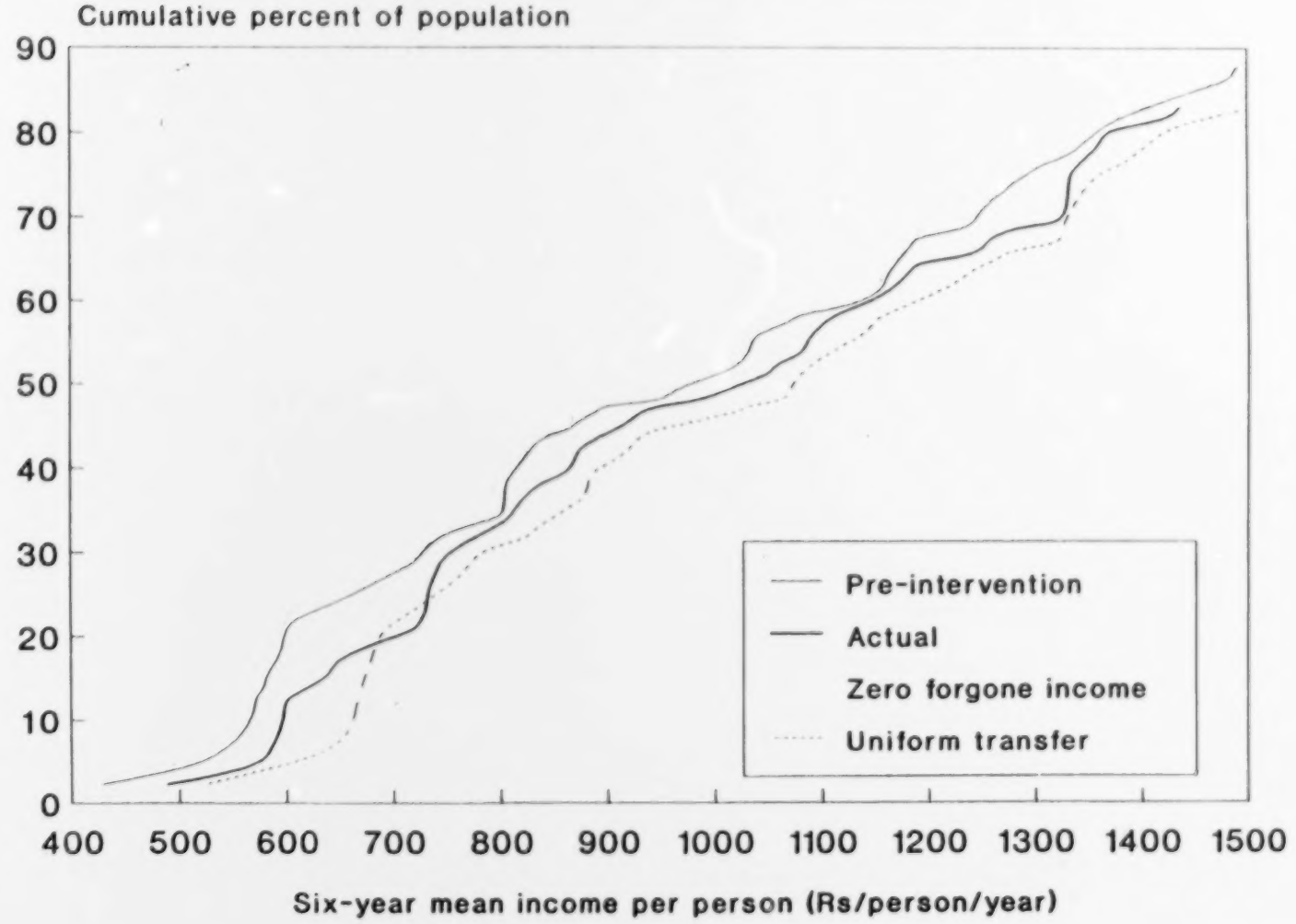
[23] Elsewhere we examine how large these costs would need to be to reverse the conclusions we come to below; we find that they would need to be equivalent to about 40% of gross disbursement (Ravallion and Datt, 1993). This would seem to be a high figure, leading us to conjecture that our qualitative conclusion could well be robust to a complete accounting of these costs.

Figure 2: Income distributions with and without earnings from public employment projects



10 Conclusions

We have proposed an empirical approach to estimating the impact on intra-household time allocation of employment on public-works projects. The model explains time allocation conditional on public-works employment, which is allowed to be endogenous, when suitable tests reject exogeneity. The statistical implementation is for two villages in the state of Maharashtra in India. Behavioral responses differ markedly between the villages. In Shirapur, wealthier households participate less in the projects, though there are signs that social stigmas and work disabilities dilute targeting performance somewhat. There is little sign of such effects in Kanzara. In both villages, employment in the projects is generally exogenous to time allocation, suggesting that the ideal embodied in the EGS of providing such work on demand is not being met. This confirms time-series evidence in Ravallion et al. (1993). The one exception to the exogeneity finding is for unemployment. This is consistent with the rationing of available public employment according to (in part) unemployment in other activities. In Shirapur, where a gender disaggregation of the model is feasible, there are signs of significant gender cross-effects in time allocation, such as through men taking up more own-farm work when women join the project sites. The projects also displace different activities for different genders: unemployment for men, leisure/domestic work for women. Our results suggest that the pecuniary opportunity cost to public-works participants is low, though this is more true of Shirapur, where participation is also higher. Overall, the projects do appear to generate sizable net income gains to participants, certainly far greater than implied by using market wages rates for similar work to value the foregone income. The transfer benefits alone led to a reduction in poverty, of almost the same magnitude as a uniform and undistorting allocation of the same gross budget.

Appendix 1: A Consistent Estimator and Exogeneity Test

It is crucial to our analysis that we can obtain consistent estimates of the parameters of the time allocation model. We have estimated CTAM using a limited information simultaneous tobit model where the endogenous variable appearing in the structural tobit model is also censored. As the properties of this estimator have not - to our knowledge - been discussed in the literature we shall outline them in detail here.

To simplify the exposition, consider the following two equation simultaneous tobit model, a linearization of (6) for n=2 and t=1,..,T:

$$L_{1t}^{*} = x_t'\beta + \gamma L_{2t} + u_t \tag{A1}$$

$$L_{2t}^{*} = x_t'\pi + v_t \tag{A2}$$

in which

$$\begin{bmatrix} u_t \\ v_t \end{bmatrix} \sim N[0, \Sigma] \tag{A3}$$

where

$$\Sigma = \begin{pmatrix} \sigma_{11} & \sigma_{12} \\ \sigma_{12} & \sigma_{22} \end{pmatrix}$$

and where x_t' is a row vector of K weakly exogenous variables. On defining

$$\begin{aligned} d_{jt} &= 1 \quad \text{if} \quad L_{jt}^{*} > 0 \\ &= 0 \quad \text{if} \quad L_{jt}^{*} \leq 0 \quad \text{for} \quad j=1,2, \end{aligned} \tag{A4}$$

the observation mechanism for the model (A1)-(A2) can be written as:

$$L_{jt} = d_{jt} L_{jt}^{*} \tag{A5}$$

Before we describe our estimator, it should be emphasized that model (A1) - (A5) is the *structural-form* model for the time allocation problem. In particular, the workfare employment equation (8) is derived as the solution of the household's optimization problem (discussed in section 2), and represents a structural equation of the CTAM. Also note that, while equations (A1) and (A2) imply a

triangular matrix of the coefficients of endogenous variables, they still constitute a simultaneous system because the covariance (σ_{12}) between u_t and v_t cannot be assumed to vanish.[24] Allowing for $\sigma_{12} \neq 0$ is appropriate insofar as (A1) and (A2) are derived from the same household optimization problem.

Our estimator is a generalization of that proposed by Smith and Blundell (1986) who allow only continuous endogenous variables in the structural limited dependent variable model. However, our generalization of the Smith-Blundell estimator (and the exogeneity test) is limited in scope. Specifically, the generalized estimator proposed below is consistent for a simultaneous tobit model with censored endogenous variables and a triangular matrix of structural coefficients. While that is appropriate here, we do not offer a consistent estimator of the general simultaneous tobit model where no restrictions are placed on the structural coefficients.[25]

Letting $\alpha v_t + \epsilon_t$ be the value of u_t conditional on v_t, the structural equation for L_{1t} is

$$L_{1t}^{*} = x_t'\beta + \gamma L_{2t} + \alpha v_t + \epsilon_t \qquad \text{(A6)}$$
$$= w_t'\delta + \epsilon_t$$

where $\epsilon_t \sim N(0,\sigma_{11.2})$, $\alpha = \sigma_{12}/\sigma_{22}$, $\sigma_{11.2} = \sigma_{11} - \sigma_{12}^2/\sigma_{22}$, $w_t' = (x_{1t}', L_{2t}, v_t)$ *and* $\delta' = (\beta', \gamma, \alpha)$

We shall refer to the model (A2), (A4)-(A6) as the "expanded CTAM", recognizing that the initial CTAM ((A1)-(A5)) has been augmented by the inclusion of v_t as an extra variable in (A6). Note first that the joint likelihood function for (L_1, L_2) can be decomposed into a conditional likelihood function for L_1, given L_2, and a marginal likelihood function for L_2. Then

$$L^{*}(L_1, L_2) = \Omega_1^{*}(L_1; \lambda^{*\prime}, \theta' | L_2).\Omega_2(L_2; \theta')$$

---

[24] If the covariance matrix was a diagonal matrix, the L_2 would cease to be endogenous in (A1), and a simple tobit ML estimator would yield consistent estimates.

[25] Amemiya (1974) proposed an estimator for such a model. But the estimation procedure only uses the subset of observations for which all the endogenous variables attain their non-limit values. For our data, this would have meant throwing out most of the sample.

for the initial CTAM, and

$$L(L_1,L_2) = \Omega_1(L_1;\lambda',\theta'|L_2).\Omega_2(L_2;\theta')$$

for the expanded CTAM, where $\lambda^{*\prime} = (\beta',\gamma,\sigma_{11},\sigma_{12})$, $\lambda' = (\beta',\gamma,\alpha,\sigma_{11.2})$, $\theta' = (\pi',\sigma_{22})$. It is easily shown that $\Omega_1^* = \Omega_1$ and hence $L^* = L$.[26]

Let us additionally define $\lambda^{0\prime} = (\beta',\gamma,\sigma_{11})$. Then, following Engle, Hendry and Richard (1983), L_2 is defined to be exogenous for $\lambda^{0\prime}$ if and only if

$$L^*(L_1,L_2) = \Omega_1^*(L_1;\lambda^{0\prime}|L_2).\Omega_2(L_2;\theta')$$

where $\lambda^{0\prime}$ and θ' are not subject to any cross restrictions.[27] A sufficient condition for $L^*(L_1,L_2)$ to be factorizable as above is that $\sigma_{12} = 0$. Given a one-to-one correspondence between $(\sigma_{11}, \sigma_{12})$ and $(\sigma_{11.2}, \alpha)$, $\sigma_{12} = 0$ is equivalent to $\alpha = 0$. When $\alpha = 0$, $\sigma_{11.2} = \sigma_{11}$, $\Omega_1^*(L_1;\lambda^{0\prime}|L_2) = \Omega_1(L_1;\lambda^{0\prime}|L_2)$ and $L(L_1,L_2)$ can be factorized as

$$L(L_1,L_2) = \Omega_1(L_1;\lambda^{0\prime}|L_2).\Omega_2(L_2;\theta')$$

Testing the exogeneity of L_{2t} for $\lambda^{0\prime}$ in equation (A1) is equivalent to testing for $\alpha = 0$.

Following Smith and Blundell (1986), our estimation procedure involves replacing v_t in the expanded CTAM with a consistent estimate $\hat{v}_t$, which yields

$$\begin{aligned} L_{1t}^* &= x_t'\beta + \gamma L_{2t} + \alpha\hat{v}_t + e_t \\ &= \hat{w}_t'\delta + e_t \end{aligned} \qquad \text{(A7)}$$

---

[26] This is proved in an addendum available from the authors.

[27] Since we have a static model, weak exogeneity coincides with strong exogeneity (Engle, Hendry and Richard 1983). We shall thus stick to the term "exogeneity".

In the Smith-Blundell model, $\hat{v}_t$ is estimated by the OLS residuals from the second (uncensored) equation. Here we propose that $\hat{v}_t$ is obtained as the residual from the tobit model for L_{2t}, which is a consistent estimator of v_t given by:

$$v_t = L_{2t} - F_{2t}x_t'\pi - \sigma_{22}f_{2t} \tag{A8}$$

where

$$F_{2t} = \frac{1}{\sqrt{2\pi\sigma_{22}}}\int_{-\infty}^{x_t'\pi} \exp\left[-\eta^2/(2\sigma_{22})\right]d\eta$$

$$f_{2t} = \frac{1}{\sqrt{2\pi\sigma_{22}}} \exp\left[-(x_t'\pi)^2/(2\sigma_{22})\right]$$

We then estimate (A7) as a standard tobit.

To derive the key properties of this estimator, we write the log-likelihood of the expanded CTAM as follows

$$\begin{aligned} \ln L = \sum_t \Bigg[& (1-d_{1t})\ln(1-F_{1t}) - \frac{d_{1t}}{2}\ln\sigma_{11.2} - \frac{d_{1t}}{2\sigma_{11.2}}\epsilon_t^2 \\ & + (1-d_{2t})\ln(1-F_{2t}) - \frac{d_{2t}}{2}\ln\sigma_{22} - \frac{d_{2t}}{2\sigma_{22}}v_t^2 \Bigg] \end{aligned} \tag{A9}$$

where

$$F_{1t} = \frac{1}{\sqrt{2\pi\sigma_{11.2}}}\int_{-\infty}^{w_t'\delta} \exp\left[-\eta^2/(2\sigma_{11.2})\right]d\eta$$

$$f_{1t} = \frac{1}{\sqrt{2\pi\sigma_{11.2}}} \exp\left[-(w_t'\delta)^2/(2\sigma_{11.2})\right]$$

Now consider the distribution of the ML estimator of $\lambda' = (\delta', \sigma_{11.2})$ given that the parameters $\theta' = (\pi', \sigma_{22})$ in the log-likelihood function have been replaced by their consistent estimates. Then, following Amemiya

(1979, equation 3.27), we can write

$$\hat{\lambda} - \lambda \overset{a}{=} -\left[E\frac{\partial^2 lnL}{\partial\lambda\partial\lambda'}\right]^{-1}\left[\frac{\partial lnL}{\partial\lambda} + E\frac{\partial^2 lnL}{\partial\lambda\partial\theta'}(\hat{\theta}-\theta)\right] \tag{A10}$$

(where $\overset{a}{=}$ denotes that each side has the same asymptotic distribution.) Given that the expectations are taken conditional on v_i, it follows from Amemiya (1973) that

$$plim\ (\hat{\lambda} - \lambda) = -\left[E\frac{\partial^2 lnL}{\partial\lambda\partial\lambda'}\right]^{-1}\left[plim\frac{\partial lnL}{\partial\lambda} + E\frac{\partial^2 lnL}{\partial\lambda\partial\theta'}plim\,(\hat{\theta}-\theta)\right] = 0 \tag{A11}$$

since $plim\ (\partial lnL/\partial\lambda)=0$ and $plim(\hat{\theta}-\theta)=0$. Thus the estimator $\hat{\lambda}$ is consistent.

Next, we derive the asymptotic covariance matrix of $\hat{\lambda}$. From (A10),

$$V(\hat{\lambda}) = -\left[E\frac{\partial^2 lnL}{\partial\lambda\partial\lambda'}\right]^{-1} + \left[E\frac{\partial^2 lnL}{\partial\lambda\partial\lambda'}\right]^{-1} E\frac{\partial^2 lnL}{\partial\lambda\partial\theta'}\,V(\hat{\theta})\,E\frac{\partial^2 lnL}{\partial\theta\partial\lambda'}\left[E\frac{\partial^2 lnL}{\partial\lambda\partial\lambda'}\right]^{-1} \tag{A12}$$

In Appendix 2 we show that this can be written as:

$$V(\hat{\lambda}) = (\mathbf{W}'\mathbf{A}\mathbf{W})^{-1} + \alpha^2(\mathbf{W}'\mathbf{A}\mathbf{W})^{-1}\left\{\mathbf{W}'\mathbf{A}\begin{bmatrix}I\\0\end{bmatrix}\mathbf{B}\mathbf{X}\,V(\hat{\theta})\,\mathbf{X}'\mathbf{B}'[I\quad 0]\mathbf{A}\mathbf{W}\right\}(\mathbf{W}'\mathbf{A}\mathbf{W})^{-1} \tag{A13}$$

where

$$V(\hat{\theta}) = -\left[E\frac{\partial^2 \ln\Omega_2}{\partial\theta\partial\theta'}\right]^{-1}$$

and the first term on the right-hand-side of (A13) is the standard covariance matrix of the (tobit) ML estimator $\hat{\lambda}$ (see Appendix 2 for details on the definition of matrices $\mathbf{W}$, $\mathbf{A}$, $\mathbf{B}$, $\mathbf{X}$). It is obvious from (A13) that under the null hypothesis of exogeneity, i.e. $\alpha=0$, the covariance matrix of $\hat{\lambda}$ collapses to the covariance matrix of the tobit ML estimator $\hat{\lambda}$. Standard t-ratios for α from a ML estimate of (A7) can thus be used to test for the exogeneity of the censored regressor L_{2i}. Even when exogeneity is rejected, the ML estimator of (A7) is consistent, with its covariance matrix then given by (A13).

To simplify the exposition, we have only considered a single censored (potentially) endogenous regressor. The results are easily generalized for a vector of censored endogenous regressors.

Appendix 2: Asymptotic Covariance Matrix for the CTAM Estimator

Equation (A12) is obtained from (A10) after dropping the terms involving the covariance $\mathbf{E}\left\{\frac{\partial lnL}{\partial \lambda}\,(\hat{\theta}-\theta)'\right\}$, since $\mathbf{E}\left\{\frac{\partial lnL}{\partial \lambda}\,(\hat{\theta}-\theta)'\right\} = \mathbf{E}\left(\frac{\partial lnL}{\partial \lambda}\right)\mathbf{E}(\hat{\theta}-\theta)' = 0$, and using the result from Amemiya (1973) that $\mathbf{E}\left[\frac{\partial lnL}{\partial \lambda}.\frac{\partial lnL}{\partial \lambda'}\right] = -\mathbf{E}\frac{\partial^2 lnL}{\partial \lambda \partial \lambda'}$. Notice that in (A12), the first term is the covariance matrix for the standard Tobit ML estimator, which following Amemiya (1973) can be written as

$$-\mathbf{E}\frac{\partial^2 \ln L}{\partial \lambda \partial \lambda'} = \left(\underline{W}' A \underline{W}\right)^{-1} \tag{A14}$$

$$\underline{W} = \begin{bmatrix} W & 0 \\ Q & l \end{bmatrix} \tag{A15}$$

and $W = (w_1, w_2, \ldots w_T)'$ with dimensions $Tx(K+2)$, Q is a $Tx(K+2)$ matrix of zeros, 0 and l are column vectors of zeros and ones,

$$A = \begin{bmatrix} A_{11} & A_{12} \\ A_{12} & A_{22} \end{bmatrix} \tag{A16}$$

and A_{ij} is a TxT diagonal matrix with diagonal elements $a_{ij}(t)$ given as in Amemiya (1979; (3.17), (3.18), (3.19)). In particular:

$$a_{11}(t) = \frac{1}{\sigma_{11.2}}\left[(w_t'\delta)f_{1t} - \frac{\sigma_{11.2}f_{1t}^2}{1-F_{1t}} - F_{1t}\right] \tag{A17}$$

$$a_{12}(t) = \frac{-1}{2\sigma_{11.2}^2}\left[(w_t'\delta)^2 f_{1t} + \sigma_{11.2}f_{1t} - \frac{\sigma_{11.2}(w_t'\delta)f_{1t}^2}{1-F_{1t}}\right] \tag{A18}$$

To evaluate

$$E\,\frac{\partial^2 \ln L}{\partial\lambda\partial\theta'} = \begin{bmatrix} E\,\dfrac{\partial^2 \ln L}{\partial\delta\partial\pi'} & E\,\dfrac{\partial^2 \ln L}{\partial\delta\partial\sigma_{22}} \\ E\,\dfrac{\partial^2 \ln L}{\partial\sigma_{11.2}\partial\pi'} & E\,\dfrac{\partial^2 \ln L}{\partial\sigma_{11.2}\sigma_{22}} \end{bmatrix} \qquad \text{(A19)}$$

the following results are used:

$$\frac{\partial e_t}{\partial\pi} = -\frac{\partial(w_t'\delta)}{\partial\pi} = \alpha x_t F_{2t}$$

$$\frac{\partial e_t}{\partial\sigma_{22}} = -\frac{\partial(w_t'\delta)}{\partial\sigma_{22}} = \alpha f_{2t}/2$$

$$\frac{\partial F_{1t}}{\partial\sigma} = f_{1t}x_t$$

$$\frac{\partial F_{1t}}{\partial\sigma_{11.2}} = -\frac{1}{2\sigma_{11.2}}(w_t'\delta)f_{1t}$$

$$\frac{\partial f_{1t}}{\partial\delta} = -\frac{1}{\sigma_{11.2}}(w_t'\delta)f_{1t}x_t$$

$$\frac{\partial f_{1t}}{\partial\sigma_{11.2}} = \left[\frac{(w_t'\delta)-\sigma_{11.2}}{2\sigma_{11.2}^2}\right]f_{1t}$$

$$\frac{\partial F_{1t}}{\partial\pi} = -\alpha x f_{1t}F_{2t}$$

$$\frac{\partial F_{1t}}{\partial\sigma_{22}} = \frac{\alpha}{2\sigma_{11.2}}(w_t'\delta)f_{1t}f_{2t}$$

$$\frac{\partial f_{1t}}{\partial\pi} = \frac{\alpha}{\sigma_{11.2}}(w_t'\delta)\,x_t f_{1t}F_{2t}$$

$$\frac{\partial f_{1t}}{\partial\sigma_{22}} = \frac{\alpha}{\sigma_{11.2}}(w_t'\delta)\,x_t f_{1t}(f_{2t}/2)$$

From the log-likelihood function (A9), we have

$$\frac{\partial \ln L}{\partial \delta} = \sum_t \left[\frac{(d_{1t} e_t w_t)}{\sigma_{11.2}} - \frac{(1-d_{1t})}{(1-F_{1t})} f_{1t} w_t \right]$$

$$\frac{\partial \ln L}{\partial \sigma_{11.2}} = \frac{1}{2\sigma_{11.2}} \sum_t \left[\frac{(1-d_{1t})}{(1-F_{1t})} (w_t'\delta) f_{1t} - d_{1t} + \frac{d_{1t} e_t^2}{\sigma_{11.2}} \right]$$

Thus, using (A17) and (A18), we can write

$$E \frac{\partial^2 \ln L}{\partial \delta \partial \pi'} = -\alpha \sum_t a_{11}(t) F_{2t} w_t x_t' = -\alpha W' A_{11} C X$$

$$E \frac{\partial^2 \ln L}{\partial \sigma_{11.2} \partial \pi'} = -\alpha \sum_t a_{12}(t) F_{2t} x_t' = -\alpha l' A_{12} C X$$

$$E \frac{\partial^2 \ln L}{\partial \delta \partial \sigma_{22}} = -\alpha \sum_t a_{11}(t) (f_{2t}/2) w_t = -\alpha W' A_{11} D l$$

$$E \frac{\partial^2 \ln L}{\partial \sigma_{11.2} \partial \sigma_{22}} = -\alpha \sum_t a_{12}(t) (f_{2t}/2) = -\alpha l' A_{12} D l$$

where C and D are diagonal matrices with F_{2t} and $f_{2t}/2$ as their t-th diagonals respectively.

Further defining $B = [C \;\; D]$ and $\underset{\sim}{X} = \begin{bmatrix} X & 0 \\ Q & l \end{bmatrix}$, we can write

$$E \frac{\partial^2 \ln L}{\partial \lambda \partial \theta'} = -\alpha \underset{\sim}{W}' A \begin{bmatrix} I \\ O \end{bmatrix} B \underset{\sim}{X}$$

and so

$$V(\hat{\lambda}) = (\underset{\sim}{W}' A \underset{\sim}{W})^{-1} + \alpha^2 (\underset{\sim}{W}' A \underset{\sim}{W})^{-1} \left\{ \underset{\sim}{W}' A \begin{bmatrix} I \\ 0 \end{bmatrix} B \underset{\sim}{X} V(\hat{\theta}) \underset{\sim}{X}' B' [I \;\; 0] A \underset{\sim}{W} \right\} (\underset{\sim}{W}' A \underset{\sim}{W})^{-1}$$

$$V(\hat{\theta}) = -\left[E \frac{\partial^2 \ln \Omega_2}{\partial \theta \partial \theta'} \right]^{-1}$$

as claimed in Appendix 1.

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Table 1: Time allocation and household attributes by days of public-works participation in Shirapur

	Days of participation per year per adult household member			
	0	< 30	30-60	> 60
Number of household years	93	67	22	16
Average days per year per adult:				
Public works	0	11.92	41.65	89.34
Wage labor	31.18	41.20	34.48	40.39
Own farm labor	49.87	62.44	48.00	39.38
Unemployment	6.82	16.64	26.57	23.94
Leisure/domestic work	212.24	181.86	160.56	140.05
Total available days	300.12	314.07	311.26	333.11
Average days per year per adult male:				
Public works	0	14.71	46.37	80.24
Wage labor	52.45	54.71	44.43	50.23
Own farm labor	65.68	80.62	68.51	58.36
Unemployment	7.72	19.72	32.89	27.07
Leisure/domestic work	162.21	136.36	128.04	117.50
Total available days	288.06	306.12	320.24	333.40
Average days per year per adult female:				
Public works	0	9.01	37.26	99.58
Wage labor	11.60	27.12	25.25	29.32
Own farm labor	35.32	43.49	28.99	18.03
Unemployment	6.00	13.44	20.71	20.42
Leisure/domestic work	258.29	229.29	190.71	165.42
Total available days	311.21	322.36	302.93	332.77
*Household attributes:**				
MRAST	60.95	47.77	36.20	17.33
CASTE	2.08	1.81	1.68	2.31
MLITM	1.33	1.21	0.89	0.90
MLITF	0.98	0.43	0.20	0.10
SCHYRH	2.23	2.82	2.23	1.25
LANDU	4.93	3.69	1.87	1.37
LANDI	0.43	0.43	0.26	0.07
BULOK	0.72	0.81	0.55	0.50
OLVSTK	2.85	2.34	1.86	0.94
ADULTM	1.99	2.18	1.73	1.69
ADULTF	2.16	2.09	1.86	1.50
CHILD	2.43	1.78	1.68	2.19
HHSIZE	6.58	6.04	5.27	5.38

* See Table 4 for definitions of the following variables.

Table 2: Time allocation and household attributes by days of public-works participation in Kanzara

	Days of participation per year per adult household member		
	0	< 30	> 30
Number of household years	137	48	13
Average days per year per adult:			
Public works	0	8.19	77.26
Wage labor	31.94	72.34	63.02
Own farm labor	61.52	42.03	12.27
Unemployment	8.52	17.36	32.38
Leisure/domestic work	230.15	184.20	150.49
Total available days	332.14	324.12	335.42
Average days per year per adult male:			
Public works	0	13.10	118.90
Wage labor	36.22	71.48	44.50
Own farm labor	109.79	65.09	15.12
Unemployment	7.69	16.03	29.25
Leisure/domestic work	192.15	166.81	151.41
Total available days	345.86	332.50	359.19
Average days per year per adult female:			
Public works	0	3.29	33.42
Wage labor	28.42	73.19	82.51
Own farm labor	21.77	18.97	9.26
Unemployment	9.21	18.70	35.69
Leisure/domestic work	261.44	201.59	149.52
Total available days	320.84	315.74	310.39
*Household attributes:**			
MRAST	64.99	15.57	12.72
CASTE	2.06	2.58	3.00
MLITM	1.37	1.22	0.90
MLITF	1.04	0.38	0.24
SCHYRH	4.30	3.17	4.69
LANDU	5.38	1.39	1.04
LANDI	0.70	0.09	0.00
BULOK	2.04	0.71	0.15
OLVSTK	2.92	1.38	1.38
ADULTM	1.74	2.31	1.54
ADULTF	2.11	2.31	1.46
CHILD	2.59	2.00	1.23
HHSIZE	6.44	6.63	4.23

* See Table 4 for definitions of the followirg variables.

Table 3: Non-parametric estimates of average number of days displaced and average foregone incomes attributed to public works.

	Shirapur						Kanzara	
	Male		Female		Total*		Total*	
Activity	Days	Value	Days	Value	Days	Value	Days	Value
Wage labor	24.02	197.78	5.32	22.97	29.34	220.75	3.05	17.56
Own farm/ other activities	-18.09	-4.00	0.26	0.06	-17.83	-3.94	45.11	4.12
Unemployment	-27.44	0	-7.20	0	-34.64	0	-4.21	0
Leisure/ domestic work	79.66	0	49.99	0	129.65	0	35.26	0
Total:	58.15	193.78	48.36	23.03	112.86	216.81	84.39	21.68
Public works	58.15	454.23	48.36	233.82	112.86	717.95	84.39	571.97

Note: Values are in 1983-84 rupees; all figures are annual, averaged over all household-years with participation.
* Totals include children.

Table 4: Explanatory variables in the time allocation model

Variable	Description	Shirapur Mean	Shirapur Std. Dev.	Kanzara Mean	Kanzara Std. Dev.
INT	intercept				
D80 - D84	dummy variables for the years 1980-81 to 1984-85	0.167	0.374	0.167	0.374
MRAST	average real assets of the household (6-year average: Rs. '000 at 1983 prices)	50.335	40.596	49.575	66.392
MRASTSQ	square of MRAST x 10^{-2}	41.734	57.258	68.433	153.79
CASTE	caste ranking of the household	1.965	1.151	2.248	0.984
MLITM	average number of literate adult males in the household (6-year average).	1.202	1.022	1.303	0.894
MLITF	average number of literate adult females in the household (6-year average)	0.636	1.216	0.828	0.954
MLITSQ	square of the average number of literate adults in the household (6-year average)	6.865	13.056	6.963	8.981
SCHYRH	years of schooling of the head of the household	2.364	3.532	4.051	4.011
LANDU	owned un-irrigated land (hectares)	3.885	3.768	4.128	5.740
LANDI	owned irrigated land (hectares)	0.384	0.816	0.503	1.235
LANDSQ	square of total owned land	33.736	50.934	63.590	166.52
BULOK	number of owned bullocks	0.707	0.990	1.591	2.244
OLVSTK	number of owned other livestock (other than bullocks)	2.424	2.320	2.444	2.597
LVSTKSQ	square of owned total livestock	17.232	26.552	34.742	61.632
ADULTM	number of adult males in the household	2.000	1.008	1.864	1.245
ADULTF	number of adult females in the household	2.051	1.107	2.116	1.279
CHILD	number of children in the household	2.101	1.414	2.359	1.852
HHSZSQ	square of household size	43.232	35.096	52.641	75.764
DISABM1	number of men (women or children)	0.505	0.682	0.268	0.498
DISABM2	reporting some work disability.	0.020	0.141	0.051	0.242
DISABF1	Type 1 is when the person cannot do	0.242	0.453	0.278	0.603
DISABF2	more than light farm or domestic work.	0.020	0.141	0.131	0.381
DISABC1	Type 2 is when the person	0.611	0.858	0.212	0.519
DISABC2	is completely disabled.	1.434	1.219	2.126	1.742
TM	total available adult male days x 10^{-1} (including leisure and domestic work but excluding sick and non-resident days)	60.119	31.062	63.840	42.463
TF	total available adult female days x 10^{-1} (including leisure and domestic work but excluding sick and non-resident days)	64.704	37.083	67.509	42.137

Table 5: Estimates of the public-works employment equations

	Shirapur				Kanzara	
	Male		Female		Male and Female	
Explanatory variable	Parameter estimate	t-ratio	Parameter estimate	t-ratio	Parameter estimate	t-ratio
INT	-239.12	-2.615	-736.26	-3.308	-133.28	-1.565
D80	-20.141	-0.805	17.063	0.543	16.733	0.421
D81	-26.289	-1.029	-38.308	-1.137	90.469	2.227
D82	-101.12	-3.511	-58.189	-1.636	44.475	1.049
D83	-105.12	-3.568	-16.409	-0.464	-76.737	-1.584
D84	-52.352	-1.865	-23.527	-0.666	-30.006	-0.616
MRAST	-4.6309	-2.596	7.9138	2.07	3.3018	0.955
MRASTSQ	3.5226	3.23	-13.887	-3.015	-2.6025	-0.855
CASTE	-27.18	-3.105	-51.278	-3.803	26.961	1.925
MLITM	19.039	0.871	38.907	0.911	-53.03	-1.227
MLITF	30.522	1.036	83.945	1.09	-68.726	-1.448
MLITSQ	-3.3167	-0.825	-35.748	-1.981	4.9971	0.454
SCHYRH	4.475	1.499	-0.0527	-0.012	18.979	3.838
LANDU	49.426	2.928	17.413	0.533	-73.487	-2.618
LANDI	49.899	2.415	11.888	0.28	-51.11	-0.741
LANDSQ	-5.0235	-3.726	-2.8734	-0.734	3.3419	1.439
BULOK	6.612	0.492	-6.2632	-0.324	-13.377	-0.524
OLVSTK	-6.5347	-0.785	-61.012	-3.867	0.50741	0.027
LVSTKSQ	0.0092	0.011	5.7802	3.643	0.007817	0.002
ADULTM	91.18	2.258	213.8	2.413	-17.223	-0.377
ADULTF	97.368	2.35	279.73	3.246	32.562	0.827
CHILD	126.46	2.73	298.02	3.656	-281.17	-0.076
HHSZSQ	-8.393	-3.615	-19.712	-3.155	0.58997	0.543
DISABM1	-13.952	-0.757	-49.5	-1.62	29.93	0.905
DISABM2	-72.772	-1.337	-145.83	-0.033	37.351	0.45
DISABF1	-46.173	-1.978	-125.27	-3.387	59.487	1.534
DISABF2	128.59	2.135	52.102	0.523	-65.058	-1.389
DISABC1	-87.171	-1.957	-100.28	-1.69	288.13	0.078
DISABC2	-36.01	-0.939	-40.941	-0.853	239.6	0.065
TM	1.8286	2.863	0.37994	0.397	1.8972	1.465
TF	0.8942	1.166	3.2874	2.908	-0.44814	-0.414
SIGMA	81.514	12.207	83.556	9.715	101.24	10.577
Log Likelihood	-541.654		-366.731		-398.818	
Correlation between actual and predicted values	0.60		0.68		0.70	

Table 6: Estimates of the conditional time allocation model for males in Shirapur

	Wage labor		Own Farm labor		Unemployment	
Explanatory variable	Parameter estimate	t-ratio	Parameter estimate	t-ratio	Parameter estimate	t-ratio
INT	-99.108	-0.938	-323.82	-3.621	-204.96	-3.638
D80	25.256	0.782	42.66	1.485	71.128	4.338
D81	24.627	0.738	81.303	2.746	102.21	6.136
D82	79.347	2.342	130.15	4.288	53.967	2.247
D83	24.945	0.705	99.986	3.158	-23.942	-0.862
D84	12.858	0.361	49.966	1.583	-19.525	-0.915
MRAST	3.7335	2.054	-1.1821	-0.708	-7.0378	-4.668
MRASTSQ	-2:507	-2.38	1.2065	1.24	5.0658	4.913
CASTE	-34.398	-3.384	3.2682	0.358	-46.441	-6.239
MLITM	31.71	1.344	-36.346	-1.914	39.972	2.749
MLITF	10.304	0.312	2.5935	0.09	38.069	2.016
MLITSQ	-5.4919	-1.301	0.63487	0.19	-5.0361	-1.949
SCHYRH	-1.4867	-0.401	-3.5673	-1.208	2.9862	1.673
LANDU	-52.107	-3.524	38.891	2.506	66.419	4.303
LANDI	-70.803	-3.195	66.298	3.149	48.337	2.813
LANDSQ	3.303	3.246	-3.1457	-2.629	-6.3599	-4.589
BULOK	-64.74	-4.342	52.133	3.895	21.095	2.534
OLVSTK	-24.507	-2.428	21.542	2.494	-3.6086	-0.661
LVSTKSQ	3.3202	3.812	-2.2711	-2.995	-0.48747	-0.896
ADULTM	90.434	2.035	51.241	1.327	81.207	2.998
ADULTF	17.307	0.354	70.91	1.657	85.698	2.618
CHILD	-56.122	-0.976	98.232	1.889	177.26	4.697
HHSZSQ	-4.1224	-1.653	-3.3555	-1.611	-7.3587	-3.581
DISABM1	-69.02	-3.076	-6.0312	-0.311	-20.228	-1.923
DISABM2	-139.97	-1.805	49.672	0.776	-104.11	-2.749
DISABF1	35.01	1.197	-94.424	-3.584	-57.159	-3.692
DISABF2	-64.349	-0.811	212.9	3.132	83.116	1.614
DISABC1	121.72	2.238	-109.02	-2.231	-127.45	-4.083
DISABC2	80.553	1.666	-54.8	-1.308	-92.082	-3.8
TM	0.22798	2.95	0.21471	3.153	0.22699	4.4
TF	0.13406	1.403	0.057071	0.753	0.043303	0.953
L_5^M			-0.40238	-2.46	-1.0604	-2.78
L_5^F	-0.43057	-2.31	0.59017	3.467	0.06896	0.798
α^M					1.3029	3.303
α^F						
SIGMA	117.77	16.532	105.41	17.86	48.308	13.894
Log Likelihood	-936.548		-1026.48		-571.629	

Table 7: Estimates of the conditional time allocation model for females in Shirapur

	Wage labor		Own farm labor		Unemployment	
Explanatory variable	Parameter estimate	t-ratio	Parameter estimate	t-ratio	Parameter estimate	t-ratio
INT	48.639	0.667	-152.94	-2.879	35.612	0.492
D80	-27.156	-1.366	17.758	1.055	57.163	2.95
D81	-1.4954	-0.073	35.757	2.053	57.774	2.905
D82	-14.044	-0.666	79.561	4.545	27.773	1.322
D83	2.2664	0.104	63.156	3.474	-21.633	-0.93
D84	2.2438	0.106	29.649	1.582	-21.935	-0.969
MRAST	-3.2763	-2.432	1.459	1.608	-2.9829	-2.238
MRASTSQ	1.257	1.225	-0.36452	-0.71	1.6043	1.847
CASTE	-8.5702	-1.385	-4.68	-0.863	-12.962	-1.885
MLITM	6.5638	0.432	-2.0677	-0.186	37.922	2.362
MLITF	17.199	0.79	-12.548	-0.769	40.885	1.8
MLITSQ	-3.2192	-1.093	1.858	0.949	-7.3481	-2.488
SCHYRH	4.8335	2.118	-0.10614	-0.061	3.8099	1.474
LANDU	10.317	0.784	-4.1356	-0.554	1.6115	0.141
LANDI	8.1079	0.445	20.626	1.88	0.29169	0.017
LANDSQ	-1.0804	-0.888	-0.68056	-1.311	-0.6115	-0.625
BULOK	-17.103	-1.248	25.975	3.419	-4.6001	-0.442
OLVSTK	-12.951	-1.291	22.963	4.543	-17.418	-2.895
LVSTKSQ	-0.10812	-0.068	-1.5353	-3.463	1.1604	2.248
ADULTM	-17.499	-0.559	-15.783	-0.695	-43.409	-1.363
ADULTF	28.147	0.895	9.939	0.4	2.17	0.063
CHILD	39.211	1.142	31.533	1.014	9.8912	0.291
HHSZSQ	-1.1988	-0.653	-0.77427	-0.655	0.72736	0.381
DISABM1	63.228	4.436	2.7525	0.24	59.893	4.058
DISABM2	-39.336	-0.715	9.9192	0.267	-197.7	-0.064
DISABF1	2.5687	0.141	-24.919	-1.666	46.458	2.417
DISABF2	20.53	0.367	62.115	1.596	-220.91	-0.072
DISABC1	-0.5155	-0.016	-39.015	-1.337	2.1516	0.067
DISABC2	-4.3616	-0.155	-9.4643	-0.368	-1.0876	-0.041
TM	0.013409	0.272	0.10494	2.649	0.027786	0.59
TF	0.037756	0.65	0.10673	2.38	0.012941	0.22
L_3^M	-0.1734	-1.592			-0.1391	-1.304
L_3^F					0.029972	0.154
α^M						
α^F					0.36158	1.655
SIGMA	61.253	14.363	62.014	18.114	54.023	11.694
Log Likelihood	-628.216		-939.085		-444.925	

Table 8: Estimates of the conditional time allocation model for Kanzara

	Wage labor		Own farm labor		Unemployment	
Explanatory variable	Parameter estimate	t-ratio	Parameter estimate	t-ratio	Parameter estimate	t-ratio
INT	290.14	4.758	-298.37	-4.363	120.63	4.126
D80	6.8942	0.216	32.641	0.944	-9.113	-0.762
D81	15.824	0.471	70.5	1.946	-33.204	-2.433
D82	-7.6023	-0.219	108.28	2.913	-89.108	-6.628
D83	-5.9376	-0.169	146.16	3.88	-101.33	-7.096
D84	-51.733	-1.416	155.42	4.041	-79.739	-5.54
MRAST	-13.477	-6.481	3.3019	1.566	-1.5976	-1.274
MRASTSQ	5.8455	5.769	-0.84208	-0.869	-2.9243	-1.255
CASTE	-23.6	-1.919	26.064	1.811	-3.8407	-0.73
MLITM	-224.73	-6.874	120.23	3.373	-47.43	-2.968
MLITF	-158.2	-4.417	87.346	2.204	-33.453	-1.676
MLITSQ	32.95	4.235	-21.348	-2.685	7.5653	1.345
SCHYRH	15.171	3.832	-12.426	-2.793	5.2555	2.192
LANDU	33.01	1.989	-18.057	-1.055	-3.1977	-0.355
LANDI	31.956	1.429	2.3592	0.113	35.143	1.34
LANDSQ	-2.9565	-3.573	-0.04002	-0.053	-0.98957	-1.115
BULOK	-1.838	-0.135	58.534	4.366	12.405	1.584
OLVSTK	7.085	0.735	15.727	1.703	-8.4409	-1.21
LVSTKSQ	-0.55587	-0.611	-1.0241	-1.428	-0.24668	-0.19
ADULTM	49.876	1.298	6.5344	0.151	6.5612	0.448
ADULTF	18.288	0.616	-52.195	-1.656	13.536	1.017
CHILD	-71.001	-0.981	9.1682	0.129	20.253	0.751
HHSZSQ	-2.3236	-3.352	0.49372	0.685	-0.30667	-0.842
DISABM1	-11.9	-0.431	-10.806	-0.35	-9.8495	-0.92
DISABM2	-26.549	-0.483	61.503	1.007	1.4514	0.046
DISABF1	7.1723	0.281	40.108	1.418	22.336	1.844
DISABF2	18.133	0.548	30.349	0.909	-7.3069	-0.488
DISABC1	167.77	2.242	14.797	0.198	6.6022	0.239
DISABC2	98.965	1.376	-3.1805	-0.045	-29.059	-1.095
TM	0.20401	1.91	0.097874	0.815	0.067902	1.471
TF	0.18921	2.284	0.27814	3.182	0.047734	1.421
$L_s^M + L_s^F$	-0.42628	-3.041	-0.42949	-2.306	-0.21102	-1.215
α					0.44076	2.248
SIGMA	111.44	17.295	129.52	18.856	39.078	16.423
Log Likelihood	-952.533		-1143.01		-698.648	

Table 9: Specification tests for public employment equations

Test (d.f.)	Shirapur: male	Shirapur: female	Kanzara
		Full sample	
Skewness (1)	0.86	19.57	4.88
Kurtosis (1)	13.69	12.38	4.35
Non-normality (2)	18.46	28.42	4.95
Heteroskedasticity (1)	10.42	3.06	16.62
Joint (3)	18.46	28.54	17.66
	Excluding three households for Shirapur and one household for Kanzara		
Skewness (1)	1.71	3.90	2.71
Kurtosis (1)	0.09	0.06	3.63
Non-normality (2)	2.58	4.59	3.65
Heteroskedasticity (1)	9.05	0.75	6.35
Joint (3)	11.74	4.73	10.76

Note: The heteroskedasticity test is a test for zero covariance between the second-moment residuals and squared predicted values of the regression function. All tests are asymptotically central χ^2. The 95 percent critical values for 1, 2, and 3 degrees of freedom are 3.84, 5.99, and 7.82; the corresponding 99 percent critical values are 6.64, 9.21, and 11.34 respectively.

Table 10: Alternative parameter estimates for public-works employment in the conditional time allocation model

Equation	Right hand side variable	Parameter estimate (t-ratio) Full sample	Pruned sample*
Shirapur: males			
Wage labor	Male public works	0	0
	Female public works	-0.431 (-2.31)	-0.394 (-1.533)
Own farm labor	Male public works	-0.402 (-2.46)	-0.332 (-1.583)
	Female public works	0.590 (3.467)	0.714 (3.077)
Unemployment	Male public works	-1.06 (-2.78)	-1.024 (-2.595)
	Male public works residual	1.303 (3.303)	1.312 (3.152)
	Female public works	0.069 (0.798)	0.045 (0.409)
Shirapur: females			
Wage labor	Male public works	-0.173 (-1.592)	-0.091 (-0.761)
	Female public works	0	0
Own farm labor	Male public works	0	0
	Female public works	0	0
Unemployment	Male public works	-0.139 (-1.304)	-0.074 (-0.542)
	Female public works	0.03 (0.154)	-0.032 (-0.142)
	Female public works residual	0.362 (1.655)	0.66 (2.279)
Kanzara:			
Wage labor	Public works	-0.426 (-3.041)	-0.195 (-1.185)
Own farm labor	Public works	-0.429 (-2.306)	-0.394 (-1.995)
Unemployment	Public works	-0.211 (-1.215)	-0.236 (-1.165)
	Public works residual	0.441 (2.248)	0.388 (1.695)

* Excluding three households for Shirapur, one household for Kanzara.

Table 11: Effects of an extra unit of public-works employment on time allocation of males and females in Shirapur.

	Extra EGS employment of:					
	Males			Females		
	Tobit coeff-icient	Prob. non-limit	Impact on time allocation	Tobit coeff-icient	Prob. non-limit	Impact on time allocation
Males:						
Wage labor	0.00	0.77	0.00	-0.43	0.77	-0.33
Own farm	-0.40	0.71	-0.29	0.59	0.71	0.42
Unemploy-ment	-1.06	0.77	-0.82	0.07	0.77	0.05
Leisure/ dom.work			0.10			-0.14
			-1.00			0.00
Females:						
Wage labor	-0.17	0.61	-0.11	0.00	0.61	0.00
Own farm	0.00	-	0.00	0.00	-	0.00
Unemploy-ment	-0.14	0.46	-0.06	0.03	0.46	0.01
Leisure/ dom.work			0.17			-1.01
			0.00			-1.00

Table 12: Effects of an extra unit of public-works employment on time allocation in Kanzara.

	Tobit coeff-icient	Prob. non-limit	Impact on time allocation
Wage labor	-0.43	0.98	-0.42
Own farm	-0.43	0.79	-0.34
Unemploy-ment	-0.21	0.92	-0.19
Leisure/ dom.work			-0.05
			-1.00

Table 13: Average number of days displaced and average foregone incomes attributed to public works.

	Shirapur						Kanzara	
	Male		Female		Total*		Total*	
Activity	Days	Value	Days	Value	Days	Value	Days	Value
Wage labor	15.55	127.96	5.30	22.85	20.85	150.81	32.96	182.31
Own farm/ other activities	-1.62	1.29	0	0	-1.62	1.29	23.04	2.01
Unemployment	36.28	0	0.19	0	36.47	0	11.70	0
Leisure/ domestic work	7.94	0	42.87	0	50.81	0	11.51	0
Total:	58.15	129.25	48.36	22.85	112.86	152.11	84.39	184.32
Public works	58.15	454.23	48.36	233.82	112.86	717.95	84.39	571.97

Note: Values are in 1983-84 rupees; all figures are annual, averaged over all household-years with participation.
* Totals include children.

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